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Monika Papież and Sławomir Śmiech and Marek A. Dąbrowski

Uniwersytet Ekonomiczny w Krakowie, Uniwersytet Ekonomiczny w Krakowie, Uniwersytet Ekonomiczny w Krakowie

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Monika Papież[1], Sławomir Śmiech[2], Marek A. Dąbrowski[3]

[1] Corresponding author: Cracow University of Economics, Rakowicka 27, 31-510 Kraków, Poland, e-mail: papiezm@uek.krakow.pl

[2] Cracow University of Economics, Rakowicka 27, 31-510 Kraków, Poland, e-mail: smiechs@uek.krakow.pl

[3] Cracow University of Economics, Rakowicka 27, 31-510 Kraków, Poland, e-mail: dabrowsm@uek.krakow.pl

Abstract

The aim of the paper is the analysis of the links between the real and financial processes in the euro area and energy and non-energy commodity prices. Monthly data spanning from 1997:1 to 2013:12 and the structural VAR model are used to analyse the relations between global commodity prices and the euro area economy. The analysis is performed for three sub-periods in order to capture potential changes in these relations in time. The main finding of the study reveals that commodity prices in the euro area do not respond to impulses from production (the economic activity), while commodity prices strongly react to impulses from financial processes, that is, the interest rates in the euro area and the dollar exchange rate to the euro (especially in the period before the global financial crisis). The study also indicates tightening the relations between energy and non-energy commodity prices.

Keywords: commodity prices, real economy, financial market, SVAR

JEL Classification:, C3, E37, E44, E47, Q17 Q43

1. Introduction

Prices of energy and non-energy sources play a key role in the economic development. High commodity prices can make production unprofitable and lead the economy to a crisis. This threat is particularly acute in countries and regions which have to import raw materials, because they do not have their own resources. The problem is deepened because commodity prices tend to co-move. Many theories and hypotheses explaining the co-movement of commodity prices have been put forward. The most general interpretation of this phenomenon is that they response to certain common and global macroeconomics shocks. The problem is that fluctuations of commodity prices are too large in comparison to what can be expected from fundamental variables. Pindyck and Rotemberg (1990) explain this excess co-movement by herd behaviours of investors, which means that the change in the overall price indexes could trigger a move in the price of any particular commodity because traders could be either in long or short position on all commodities. Lescaroux (2009) extends the studies by taking into account inventory levels (which are available for investigated oil and metal prices). Frankel and Rose (2010) refer to four plausible theories of commodity price co-movement, which cover: a strong global growth, especially in China and India (oil prices are studied in this context by Kilian, 2008, 2009a, 2009b); Hamilton, 2009; Kaufmann, 2011), easy monetary policy and low real interest rates (Frankel, 2008; Kilian, 2010), a speculation (examined by Davidson, 2008; Krugman, 2008; Śmiech and Papież, 2013; Parsons, 2010) and the risk resulting from potential geopolitical uncertainties. Sari et al. (2010) point out that oil and precious metals are denominated in US dollars, and thus co-move. A negative relationship between the value of the dollar and dollar denominated commodities follows from the law of one price for tradable goods. Another reason appears during an expected inflation period, when investors prefer to stay long in raw materials rather than assets. Similar conclusion came from Akram (2009), who also finds evidence that weaker dollar leads to higher commodity prices, and that the reduction of interest rates causes an excessive increase in prices of oil and industrial raw materials. Different results are obtained by Frankel and Rose (2010) and Alquist et al. (2011), as they do not find statistically significant relationships between real interest rates and oil prices.

The aim of this study is to examine whether commodity prices are related to real and financial processes in the Euro area macroeconomy. There are several reasons for choosing the euro area economy. Firstly, the euro area is the second largest economy worldwide, with GDP of 12,715,823 million USD in 2013 (GDP in the USA in the same year equalled 16,803,000 M.$). Secondly, euro area countries do not have sufficient supplies of energy resources, and that is why the share of import of fossil fuels in the total energy consumption is so high (60% for the EU in 2009) and still growing. Thirdly, to the best of the authors knowledge, the relations between the Euro area macroeconomy and commodity prices have not been addressed in any other study.

The analysis is based on monthly data from the period from 1997:1 to 2013:12. The real processes of the euro area are represented by the industrial production, and the financial conditions are represented by the interest rates. Bearing in mind a significant role of the American dollar, we took into account the dollar exchange rate to the euro. Commodity prices are represented by the energy price index and the non-energy price index published by the World Bank database. The structural VAR model is used to investigate the relations, as it allows for their interpretation in economic terms. In order to identify structural shocks we use – following Akram (2009) – a standard recursive structure obtained by a Choleski decomposition. The analysis is conducted in three separate sub-periods: the first one covers the time 1997:1-2002:12, the second one - 2003:1-2008:12, and the third one – 2009:1-2013:912. The division of the whole sample period into three sub-periods allows for the verification of the stability of the relations investigated and the influence of particular areas of euro economy on commodity prices. Additionally, such division allows for checking the changes in relations between energy and non-energy commodity prices, taking into consideration a growing share of biofuels (which are and element of the non-energy index) in energy consumption in the euro area.

The main outcome of the study is the finding that commodity prices in the euro area do not respond to impulses from production (economic activity). At the same time, commodity prices strongly react to impulses from financial processes, that is, the interest rates in the euro area and the dollar exchange rate to the euro (especially in the period before the global financial crisis). The study also reveals tightening the relations between energy and non-energy commodity prices.

The rest of the paper is organized as follows. Section 2 briefly presents the methodology used in the study, the description of the data can be found in Section 3, and Section 4 contains the empirical results. Final conclusions are presented in the last section.

2. Methodology

The empirical analysis is based on structural vector autoregression (SVAR) models proposed by Sims (1980). Two types of SVAR models are developed. The first one comes from Blanchared-Quah (1989) (BQ) model and assumes long-term restrictions to model innovations using the economic theory. The second one is called *AB* model (Breitung et al., 2004) and deals with short-term restrictions. In the study the latter is used:

$$Ay_t = A_1 y_{t-1} + A_2 y_{t-2} + ... + A_p y_{t-p} + B\varepsilon_t \qquad (1)$$

where: y_t contains the vector of variables, $\varepsilon_t \sim (0, I_k)$, A is $k \times k$ invertible matrix of structural coefficients, which describes contemporaneous relationship among the variables in y_t, A_i $(i = 1,2,..., p)$ are $k \times k$ coefficient matrices describing dynamic interactions between the k-variables, and B is ($k \times k$) matrix of structural coefficients representing the effects of k structural shocks. The reduced form of equation (1) can be obtained by premultiplying with the inverse of A:

$$y_t = A*_1\, y_{t-1} + A*_2\, y_{t-2} + ... + A*_p\, y_{t-p} + u_t \qquad (2)$$

where: $A*_i = A^{-1}A_i$. $u_t = A^{-1}B\varepsilon_t$, and $u_t \sim (0, \Sigma_u)$ is the symmetric variance-covariance matrix of the reduced form consisting of $k(k+1)/2$ elements.

Without making reference to a specific economic structure, the reduced form model is difficult to understand, and its parameters have no economic interpretations. In case of structural model, identification focuses on the (orthogonal) errors of the system, which are interpreted as exogenous shocks. Structural VAR model can be estimated on the basis of the reduced form model (2), which has, however, fewer parameters. Thus, in order to identify model (1), at least $k^2 + k(k-1)/2$ restrictions of the matrices *A* and *B* have to be imposed (Breitung et al., 2004). Therefore, most applications consider special cases with $A = I_k$ (B models) or $B = I_k$ (A-models). Necessary restrictions can be obtained from the economic theory, or some atheoretical rules, like a "timing scheme" for shocks proposed by Sims (1980).

To analyze the dynamic interactions between the endogenous variables of VAR(p) models, impulse response analysis can be used. Assuming that model (1) represents stationary (I(0)) process y_t, it has a Wold's moving average (MA) representation:

$$y_t = \Phi_0 u_t + \Phi_1 u_{t-1} + \Phi_2 u_{t-2} + ..., \qquad (3)$$

where $\Phi_0 = I_K$ and Φ_s are computed recursively. The (i, j) element of the matrix Φ_s, considered as a function of *s*, measures the expected response of $y_{i,t+s}$ to a unit change in innovations $u_{j,t}$.

Another useful interpretation of SVAR model is forecast error variance decompositions. To obtain it, it is enough to notice, using (3), that forecast variance y_{T+s} is expressed as:

$$\Sigma_s = Var(y_{T+s}) = \sum_{m=0}^{s-1} \Phi_m \Phi'_m \qquad (4)$$

The diagonal element of Σ_s describes variance of forecast error as a sum of errors resulting from individual structural shocks.

3. Data

To explore the relationships between commodity prices, real economy and financial indicators in the euro area, we use monthly data from the period 1997:1 to 2013:12. The analysis is based on 5 series of variables. The first one is the industrial production index (IP) in the euro area, which describes real economy in Europe. The second one is the 3-month interest rate in the euro area (IR), which describes financial economy. The data for both variables are taken from Eurostat database. The next variable is the real exchange rate (REX). The remaining two variables are the commodity price indexes, that is the energy price index (PEN) and the non-energy price index (PNEN). The data for these variables are taken from the World Bank database. The energy price index (world trade-base weights) consists of crude oil (84.6%), natural gas (10.8%) and coal (4.6%). The non-energy price index consists of metals (31.6%), fertilizers (3.6%), and agriculture (64.8%). A detailed description of variables is provided in Table 1, and basic descriptive statistics can be found in Table 2.

Next, all series are expressed as indices with their average values in 2010 equalling 100, seasonally adjusted and specified in natural logarithms.

Table 1. The dataset – Variable description.

Variable	Full name	Description	Source
IP	The industrial production index in the euro area	Euro area 17 (fixed composition) - Industrial Production Index, Total Industry (excluding construction) - NACE Rev2; Eurostat; Working day and seasonally adjusted	Eurostat
IR	The 3-month interest rate in the euro area	Nominal interest rate (NIR) minus HICP inflation: $100[\ln(1+NIR_t)-(\ln HICP_t - \ln HICP_{t-12})]$	Eurostat
REX	The real exchange rate	Index of nominal exchange rate (end of month), NER, adjusted by consumer price indexes in the US and euro area: $100 \cdot NER \cdot CPI_{US} / HICP_{EA}$, 2010=100	Eurostat, Federal Reserve Bank of St. Louis
PEN	The energy price index	Monthly index based on nominal US dollars deflated with CPIUS, 2010=100	World Bank, Federal Reserve Bank of St. Louis
PNEN	The non-energy price index	Monthly index based on nominal US dollars deflated with CPIUS, 2010=100	World Bank, Federal Reserve Bank of St. Louis

Table 2. Descriptive statistics.

	IP	IR	REX	PEN	PNEN
Mean	100.60	0.73	109.60	75.20	77.65
Median	100.46	0.93	104.84	70.01	73.38
Maximum	114.68	2.99	152.58	175.64	127.20
Minimum	87.45	-2.32	83.77	20.16	47.54
Std. Dev.	5.79	1.42	16.34	37.05	22.05
Skewness	0.29	-0.31	0.99	0.35	0.47
Kurtosis	2.97	2.01	3.14	2.05	1.99

The whole sample period is divided into three sub-periods: 1997:1-2002:12, 2003:1-2008:12, and 2009:3-2013:12. The first sub-period from January 1997 to December 2002 contains 72 observations. The industrial production index is the lowest in this period, which indicates the lowest activity of the euro area economy (the mean value of this index is 96.41). The prices of energy and non-energy sources are also the lowest in this period, while interest rates are the highest with the mean 1.99, and the median 2.20. In the first sub-period the energy price index increases by 5.5 %, while the non-energy price index decreases by 29.2%.

The second sub-period form January 2003 to December 2008 contains 72 observations. The energy price index increases at that time by 28.5%, and non-energy price index increases by 32.2%. The euro zone economy is the most active in this period, which is indicated by the highest values of the industrial production index (the mean value of this index is 105.68). The interest rates are lower than before with the mean 0.72 and the median 0.36.

In the last sub-period, from January 2009 to September 2013, the mean and the median of the real interest rates are negative, and equalled, respectively, -0.84 and -1.05. The industrial production index displays only slightly higher values than in the first sub-period, which are, however, much lower than in the second sub-period (the mean value of this index is 99.55). In the last period the energy price index increases by 78.4%, while non-energy price index increases by 16.5%.

Fig. 1. presents commodity prices and real economy and financial indicators in the Euro area.

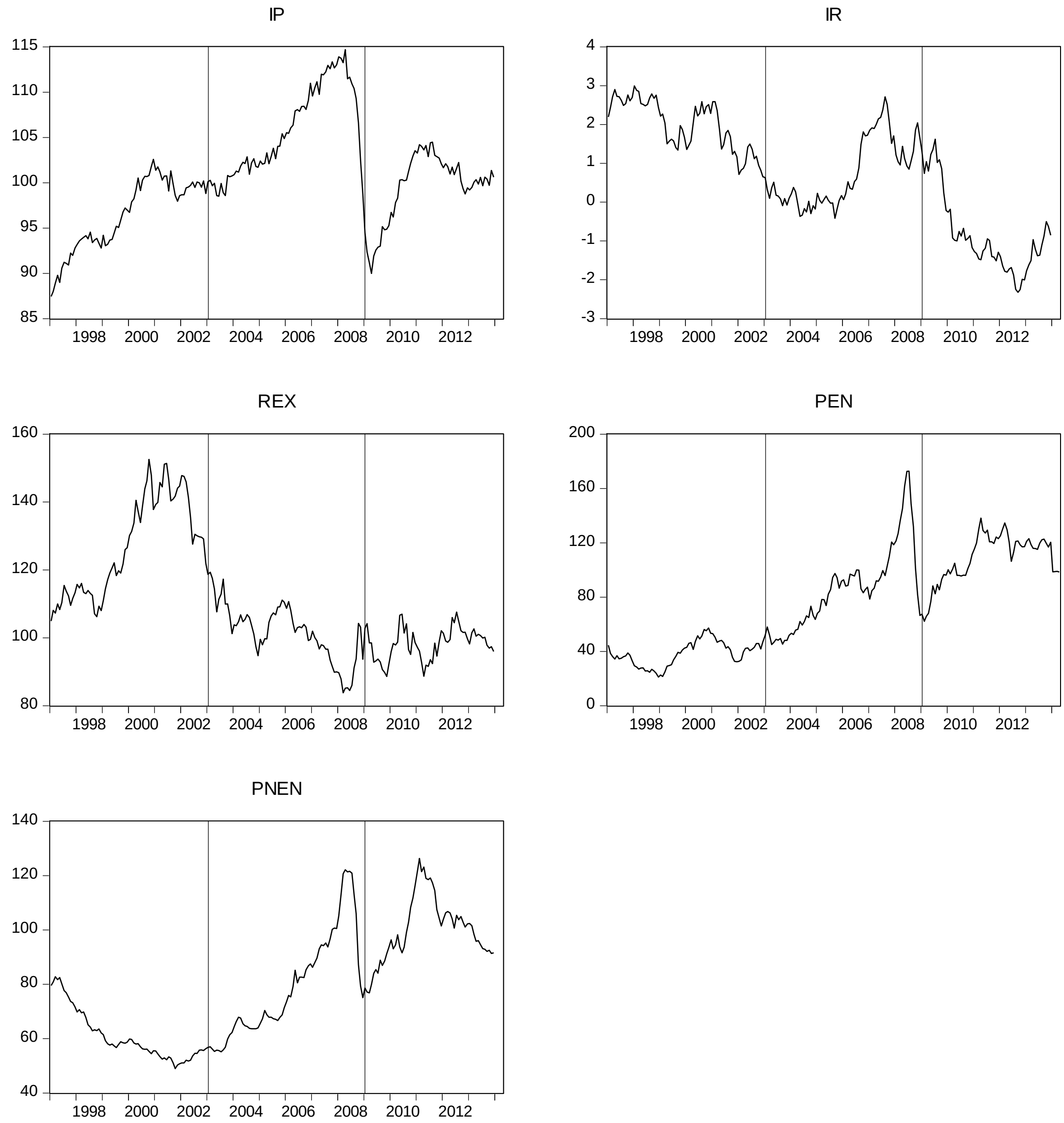

Fig. 1. Commodity prices and real economy and financial indicators in the Euro area.

4. Empirical results

4.1. Time series properties of the data

A preliminary analysis of the series is carried out before estimating the main model. The standard augmented Dickey–Fuller (ADF) (Dickey and Fuller, 1979)) unit root tests for both the intercept and the trends specifications demonstrate that all variables have unit roots for each analysed sub-period. The number of lags in the test is established using AIC criterion. The test results are presented in Table 3. The ADF unit root test confirmed that all variables were integrated of order one, i.e., I(1), thus making the test for cointegration justified.

Next, the presence of long-term relationship between integrated variables is investigated. The trace test statistic proposed by Johansen and Juselius (1990) is used. If the variables are co-integrated, the VAR in first difference would not be correctly specified, and the long-term result would be very helpful in exploring the efficient parameters of short-term dynamics. Table 4 presents the results of Johanson co-integration test. Johanson maximum likelihood approach was used to test cointegration, and it employed both maximum eigenvalue and trace statistics. According to trace test statistics and maximum eigenvalue test, there is no cointegration at 5 percent level in the first and third sub-periods. Test results demonstrate some evidence of the presence of cointegration only in the second sub-period. Trace test indicates one cointegrating equation at the 0.05 level. In contrast, maximum eigenvalue test indicates no cointegration at the 0.05 level[1]. Since the results of cointegration tests are at best ambiguous (if not suggesting the lack of cointegration), and the variables used are I(1), we use a VAR for the first differences in our five variables.

For each sub-period a number of lags for VAR was established using AIC criterion - the lag length is one for the first and second sub-periods and two for the third sub-period.

Table 3. Unit root tests results for each sub-period.

Sub-period		Level		First difference	
		intercept	intercept and trend	intercept	intercept and trend
1997:1 – 2002:12	IP	-2.7232	-1.6266	-11.6381***	-12.1336***
	IR	-0.6207	-2.5671	-7.3601***	-7.3496***
	REX	-1.4755	-0.1782	-7.0423***	-7.2472***
	PEN	-1.3676	-1.5073	-6.0474***	-5.9651***
	PNEN	-0.9121	-1.3533	-7.4002***	-7.3506***
2003:1 – 2008:12	IP	-1.1568	-1.0029	-11.9064***	-11.9143***
	IR	-1.0186	-1.8155	-7.4128***	-7.3509***
	REX	-2.3057	-2.4251	-8.4422***	-8.4539***
	PEN	-1.1136	-2.0591	-8.4195***	-8.3045***
	PNEN	-0.5708	-2.7851	-8.6750***	-8.5870***
2009:1 – 2013:12	IP	-2.3250	-2.3984	-4.1896***	-4.0177***
	IR	-2.4061	-1.0401	-6.9393***	-7.4423***
	REX	-2.5147	-2.7323	-8.5349***	-8.4427***
	PEN	-2.2888	-1.3825	-12.8698***	-13.1621***
	PNEN	-2.2941	-0.5975	-11.2026***	-12.2130***

Note: All variables in natural logs, lag lengths are determined via AIC, (***) indicates the rejection of unit root at 1 percent.

[1] Since the length of the sample is not long, and there are four series in a vector of interests, a Monte Carlo experiment is performed and the empirical critical values of trace test are determined. We find that in such case null hypothesis of no cointegration is rejected too often.

Table 4. Test for cointegration (with intercept in the CE) for each sub-period.

Sub-period	Hypothesized no. of CE(s)	Trace statistic		Max-Eigen statistic	
		Test Statistic	Critical value 0.05	Test Statistic	Critical value 0.05
1997:1 – 2002:12	None	68.040	69.819	26.663	33.877
	At most 1	41.377	47.856	19.853	27.584
	At most 2	21.523	29.797	16.327	21.132
	At most 3	5.196	15.495	3.927	14.265
	At most 4	1.269	3.841	1.269	3.841
2003:1 – 2008:12	None	70.654**	69.819	33.429	33.877
	At most 1	37.224	47.856	19.530	27.584
	At most 2	17.694	29.797	9.473	21.132
	At most 3	8.221	15.495	7.321	14.265
	At most 4	0.899	3.841	0.899	3.841
2009:1 – 2013:12	None	69.481	69.819	27.062	33.877
	At most 1	42.418	47.856	21.293	27.584
	At most 2	21.125	29.797	12.239	21.132
	At most 3	8.886	15.495	8.740	14.265
	At most 4	0.147	3.841	0.147	3.841

4.2. Structural impulse response analysis

In order to identify SVAR model, we use the Choleski decomposition of the reduced form and assume that A is an identity matrix, while B is a lower triangular matrix. To identify the shocks, we order the variables in the VAR models, and, thereby, the corresponding shocks, as $(\Delta IP, \Delta IR, \Delta REX, \Delta PEN, \Delta PNEN)'$

$$B\varepsilon = \begin{bmatrix} * & 0 & 0 & 0 & 0 \\ * & * & 0 & 0 & 0 \\ * & * & * & 0 & 0 \\ * & * & * & * & 0 \\ * & * & * & * & * \end{bmatrix} \begin{bmatrix} \varepsilon_{IP} \\ \varepsilon_{IR} \\ \varepsilon_{REX} \\ \varepsilon_{PEN} \\ \varepsilon_{PNEN} \end{bmatrix}$$

where: B is a lower diagonal matrix consistent with the Choleski decomposition, the "*" entries in the matrix represent unrestricted parameter values, and the zeros suggest that the associated fundamental shock does not contemporaneously affect the corresponding endogenous variable.

A five-variable VAR is estimated with changes in the industrial production (ΔIP), the real interest rate (ΔIR), the real exchange rate (ΔREX), the real energy prices index (ΔPEN) and the real non-energy price index (ΔPNEN). The ordering of variables is implied by the objective of this study and by the economic theory. Thus, in order to allow for reactions of commodity prices to all other variables, which is the subject matter of this paper, they are placed at the end. A similar ordering is also applied by Akram (2009). The industrial production is supposed to be the least responsive variable (its adjustment to shocks is sluggish), and that is why it is the first variable in the VAR. The ordering of the interest rate before the exchange rate is of secondary importance since the focus falls on commodity prices. Nevertheless, it is in line with the ordering applied by Arora and Tanner (2013).

The impulse-responses results for structural one standard deviation innovations in the industrial production index, the real interest rate, the real exchange rate, the energy price index and the non-energy price index are illustrated in Fig. 1 – 3 for each sub-period respectively. For instance, the impulse response of each variable in the system to an innovation in the industrial production index in the first sub-period is shown in the first column of Fig. 1 with a solid line. The dashed lines correspond to plus/minus two standard errors around the impulse responses.

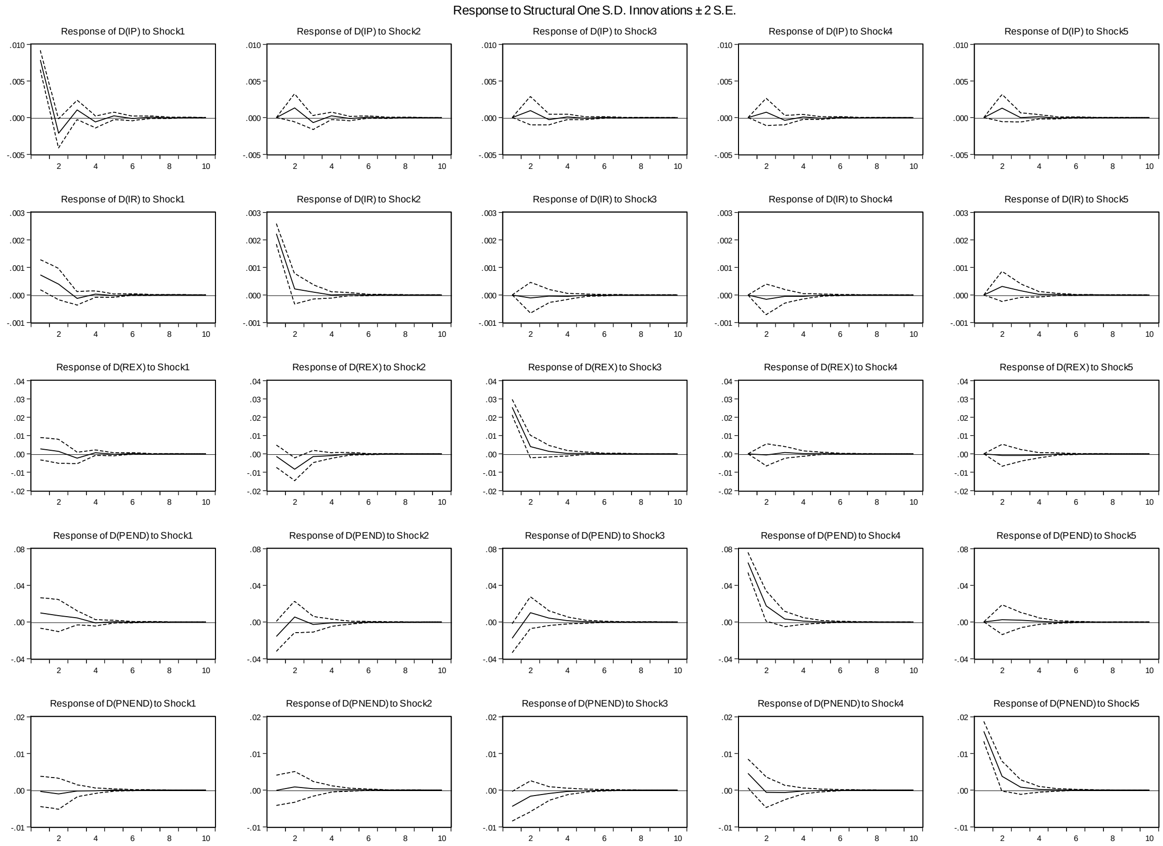

Fig. 1. The impulse-responses results in the sub-period 1997:1-2002:12.

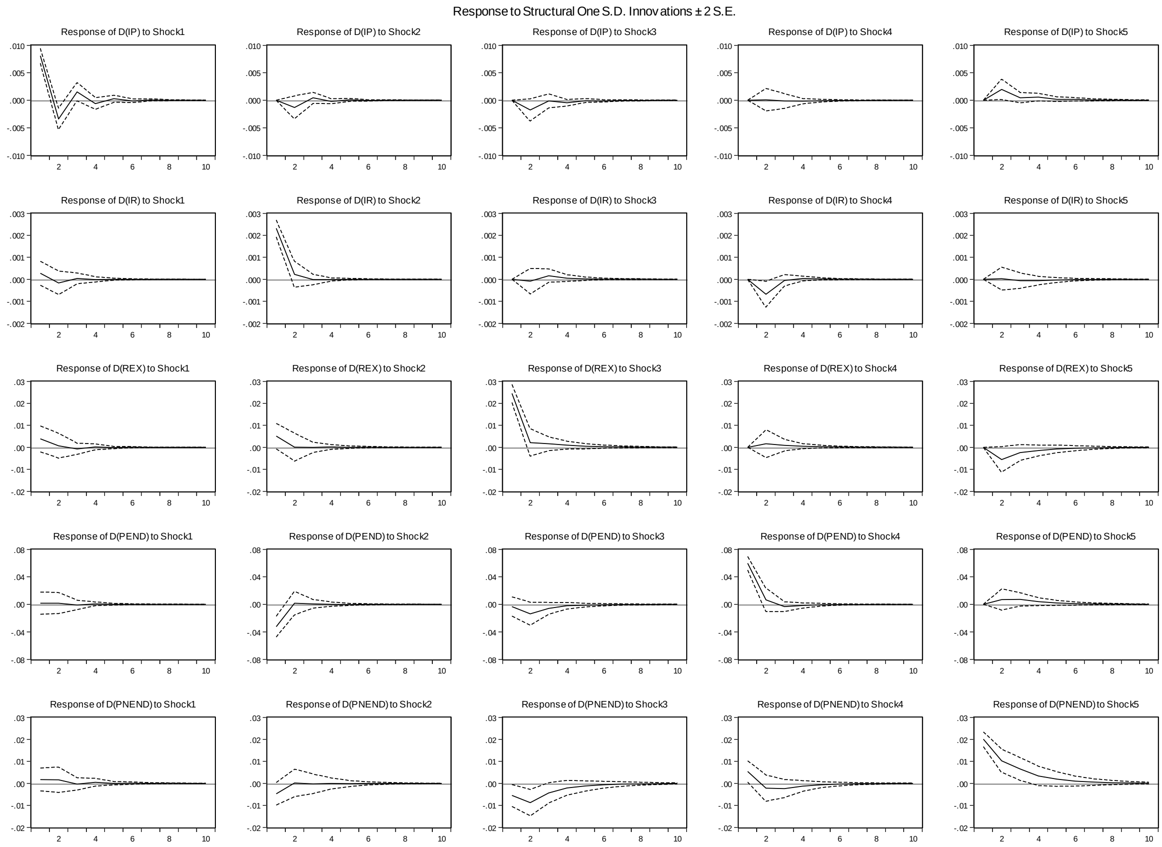

Fig. 3. The impulse-responses results in the sub-period 2003:1 – 2008:12.

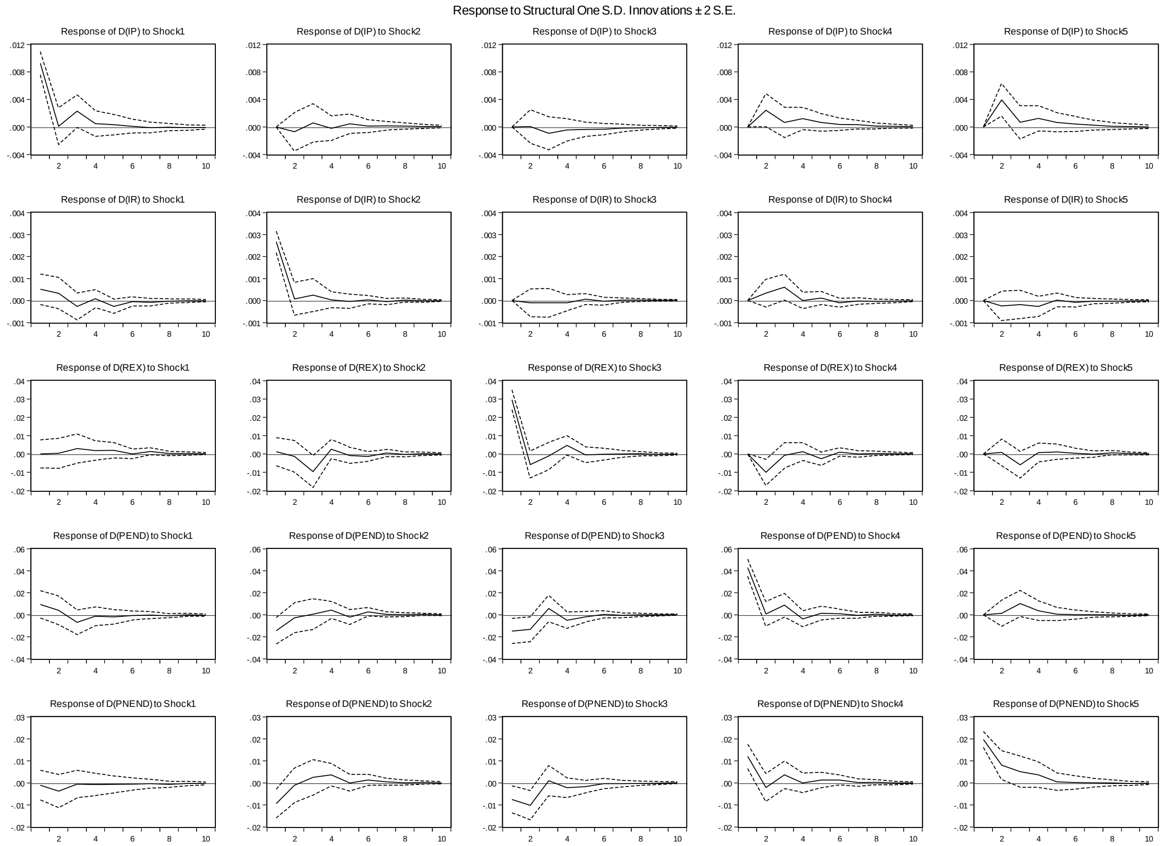

Fig. 4. The impulse-responses results in the sub-period 2009:1 - 2013:12.

In the sub-period 1997:1-2002:12 output shocks were rather neutral for all other variables except the real interest rate. The response of the latter was consistent with intuition: as economy expanded due to positive production shock, the rate of interest increased. It could also be coupled with changes in the policy rate: after all, a positive shock could trigger inflation.

Energy price index reacted negatively to shocks in the real interest rate. This is in line with Hotteling's rule, which states that the gain from storing a commodity should be equal to the interest rate. The gain includes a revaluation gain and a convenience yield and is adjusted downwards by storage cost and risk premium (see Frankel and Rose, 2010 or Śmiech et al., 2014). Such a reaction of energy prices to the interest rate make them similar to asset prices (Svensson, 2008).

The US dollar real depreciation exerts a positive impact on both commodity prices. Such a link was identified by other researchers as well (see e.g. Akram, 2009). It could be explained as follows: commodity prices are quoted in US dollars. As the US dollar depreciates, commodity prices become lower when they are expressed in other currencies. Thus, the demand for commodities increases, which results in a higher dollar price of commodities and some reversal of the initial depreciation of the US dollar.

The reaction of non-energy price index to shocks in energy prices is a significant one. It is positive, thus both commodities could be seen as related to one another. In other words and less formally, non-energy prices could not deviate too much from energy prices.

Fig. 3 illustrates impulse response functions for the sub-period 2003:1-2008:12. When they are compared with responses in the first sub-period, two differences are visible. First, interest rate shocks are much more important not only for energy prices but also for non-energy prices. It seems that commodity prices became more like price assets in the run-up to the global financial crisis. At the same time, a link running from real exchange to the energy prices rate and from industrial production to the real interest rate ceased to be significant, suggesting that the financial processes decoupled from the real economy. Second, the response of exchange rate to shocks in the interest rate is positive. It looks as an anomaly: a higher interest rate in the euro area should make the euro stronger and the US dollar weaker (a negative response), whereas the response functions suggests the opposite.

The last sub-period of 2009:1-2013:12 is a mixture of the previous two sub-periods. The interest rate is still unrelated to output shocks like in the middle sub-period, but there is no anomalous behaviour of the real exchange rate to interest rate shocks. Both commodity prices respond to interest rate shocks and exchange rate shocks as suggested by the theory. It is interesting to observe that the reaction of non-energy prices to shocks in energy prices is stronger than in the other two sub-periods.

4.3. Variance decomposition

Forecast error variance decompositions of changes in commodity price indexes at three time horizons (1, 3, 6 and 12 months) and across three sub-periods are presented in Tables 5a and 5b.[2] For both commodity price indexes own shocks account for 50-85 per cent of the forecast error variance though their contribution decrease over time.

Table 5a. Variance decomposition of the energy price index ΔPEN for each sub-period.

Sub-period	Month	Shock in:				
		ΔIP	ΔIR	ΔREX	ΔPEN	ΔPNEN
1997:1 – 2002:12	1	1.99	5.01	6.39	86.61	0.00
	3	3.05	5.16	7.92	83.68	0.18
	6	3.06	5.18	7.96	83.60	0.20
	12	3.06	5.18	7.96	83.59	0.20
2003:1 – 2008:12	1	0.06	22.88	0.21	76.85	0.00
	3	0.14	21.20	4.63	72.09	1.93
	6	0.15	21.08	4.73	71.74	2.31
	12	0.15	21.07	4.73	71.73	2.32
2009:1 – 2013:12	1	3.78	8.97	9.30	77.96	0.00
	3	5.41	7.76	15.07	67.93	3.82
	6	5.44	8.56	15.55	66.26	4.18
	12	5.48	8.56	15.55	66.23	4.19

Table 5b. Variance decomposition of the non-energy price index ΔPNEN for each sub-period.

Sub-period	Month	Shock in:				
		ΔIP	ΔIR	ΔREX	ΔPEN	ΔPNEN
1997:1 – 2002:12	1	0.05	0.00	6.54	6.96	86.44
	3	0.39	0.28	7.26	6.76	85.31
	6	0.40	0.32	7.30	6.77	85.21
	12	0.40	0.32	7.30	6.77	85.21
2003:1 – 2008:12	1	0.61	4.62	6.22	5.95	82.59
	3	0.77	3.04	16.98	5.30	73.91
	6	0.77	2.95	17.23	5.40	73.65
	12	0.77	2.94	17.24	5.40	73.64
2009:1 – 2013:12	1	0.16	12.89	8.22	21.30	57.44
	3	1.70	10.38	17.52	17.79	52.61
	6	1.76	11.54	17.57	17.36	51.77
	12	1.84	11.55	17.56	17.34	51.70

[2] Results for longer time horizons do not differ from those for 12-month horizons. Variance decompositions for (changes in) industrial production, the interest rate and the real exchange rare are available upon request.

Interest rate and exchange rate shocks make higher contributions to fluctuations in the energy price index than output shocks in the first sub-period. This dominance is even stronger in the middle sub-period when the link to interest rate clearly prevails. The contributions are more balanced in the last sub-period and a slight dominance of exchange rate shocks.

Forecast error variance decompositions of changes in non-energy price index show that the importance of all shocks has increased over time. This is especially true for interest rate shocks: their contribution was initially less than one per cent and has increased to more than 10 per cent.

One more observation refers to the fact that links between commodity prices have become tighter over time: non-energy price shocks' contribution to variance of energy price index has increased from virtually zero to 4 percent. The contribution of energy price shocks to variance of non-energy price index has increased from 7 to 17 percent. This tighter link can possibly be connected with the rising importance of biofuels in the non-energy price index (see Demirbas, 2011) or/and heightened interest of investors in financial markets of non-energy commodities. The former makes non-energy commodities similar to energy commodities, while the latter makes them similar to financial assets.

Conclusion

The euro area is a large open economy and, therefore, its real and financial developments can potentially exert certain impact on commodity prices. In order to check whether such relations hold, the structural VAR model for three sub-periods has been used. The main findings are threefold. First, economic activity in the euro area turned out to be rather neutral for commodity prices.

Second, this cannot be said about the real interest rate and the exchange rate. Energy and non-energy prices respond to shocks in the real interest rate in all sub-periods, and the link has become particularly strong in the run-up to the global financial crisis. Real exchange rate shocks have gained in their importance after the initial stage of the crisis, i.e. in the last sub-period. Even though these relations are in line with the standard commodity price determination model (see e.g. Frankel and Rose, 2010 or Śmiech et al., 2014), the model itself does not imply that financial factors should dominate over real processes. This, however, was the case in the run-up to the global financial crisis.

Third, interrelations between energy and non-energy commodity prices have become stronger over time, particularly the one running from energy to non-energy commodity prices.

It seems that non-energy commodities have become similar to financial assets: they are more sensitive to changes in the interest rate and energy prices.

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